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Changes in job stability – evidence
from lifetime job histories

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Abstract

We use lifetime job histories from the pension records to evaluate changes in job stability in Finland between 1963 and 2004. We specify a duration model and estimate the effects of elapsed duration, age, and calendar time on the hazard of job ending using individual-level panel data spanning over four decades. We find that this hazard increased during the recession years in the early 1990s but has now returned to the level that prevailed in the 1970s. We also demonstrate that the fluctuations in the hazard rate together with the changes in labor market entry rates have complicated dynamic effects on the tenure distribution, and that analysing the changes in job stability based on the elapsed duration of ongoing jobs may be quite misleading.

Key words: job stability, duration models

JEL classification numbers: J62, J63

Tiivistelmä

Tutkimuksessa arvioidaan työsuhteiden keston muutoksia vuosien 1963 ja 2004 välillä eläkerekisterien tietojen avulla. Estimoimme 42 vuoden pituisesta yksilötason paneeliaineistosta duraatiomallin, jossa työsuhteen päättymisen ja työpaikan vaihdon riskiä selitetään työsuhteen siihenastisella kestolla, työntekijän iällä ja kalenterivuodella. Tulosten mukaan työpaikan päättymisriski kasvoi 1990-luvun laman aikana, mutta on sen jälkeen palannut 1970-luvun tasolle. Osoitamme myös, että työpaikan päättymisriskin muutokset ja muutoksen uusien työpaikkojen syntymisvauhdissa aikaansaavat monimutkaisia dynaamisia vaikutuksia työsuhteitten kestojakamaan ja että siksi työsuhteiden vakauden arviointi työsuhteiden keston muutosten avulla voi olla harhaanjohtavaa.

Avainsanat: työsuhteiden kesto, duraatiomalli

JEL-luokittelu: J62, J63

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1. Introduction

There is now a relatively large literature on the changes in job stability and on the incidence of lifetime jobs in the economy. A general conclusion from the existing studies is that changes in job stability were small up to the mid 1990s (e.g. Neumark 2000, Burgess & Rees 1996), but during the more recent years job stability has declined (e.g. Farber 2007, Gregg & Wadsworth 2002). While the majority of existing studies are based on the US data, there are also several studies describing the trends in job durations in European countries. A cross-country study by OECD (2007) is probably the most recent source of comparative data across countries. We use data from Finland that could be considered as a typical European country in terms of job stability. According to the Eurostat Labor Force survey, job stability, measured by the average job tenure, was 9.9 years in Finland in 2005 which was only slightly higher than the unweighted EU average of 9.7 years.

A common problem in most existing studies of job stability is that genuine panel data where job spells could be followed over their entire span is generally lacking. This prevents modelling job durations using standard methods of survival analysis. Instead, researchers have examined the changes in the distribution of elapsed duration of ongoing jobs or used strategies based on inferring job durations from retention rates calculated from short panels or synthetic cohort data. In many ways available data on job durations resembles data that was available for research on unemployment duration in the 1980s.

In this paper we use administrative data from the Finnish pension insurance companies. These data cover a time period starting from 1963, the year following the creation of the current earnings related pension system in 1962, up to the recent pension reform in the end of 2004. We have access to a representative sample of fifteen cohorts of Finns who were employed in the private sector at some point during this 42-year interval. The oldest cohort was born in 1905 and the youngest cohort in 1975. The data include the starting and ending dates of all insured employment spells of all individuals in the sample. This allows us to create a sample of job spells that were ongoing in January 1st, 1963 or started sometime after that date and follow these spells until they end - even if the jobs last for several decades. For some cohorts this implies that we can observe all employment spells during the entire career and directly observe the completed durations of all employment spells.

We start by taking snapshots of data at regular intervals and by computing the distribution of the elapsed duration of ongoing jobs at various points in time. Since we are using administrative data, many problems related to recall errors and consistency of measurement over time can be avoided. Also problems due to non-response or panel attrition do not arise with administrative records. To verify

that administrative data is consistent with typical survey data, we compare the results from administrative records to the data on tenure distribution in the Labor Force Survey. In Finland this can be done starting from 1982. We then proceed to the analysis of differences in elapsed tenure at a given age between successive cohorts. Though this approach is commonly used in analysing changes in job stability, we would argue that it is not a particularly appealing way of analysing the *changes* in job stability. To demonstrate this, we simulate the effects of changes in the entry rate of new jobs and the hazard of job ending this on commonly used measure for job stability.

Since we have access to genuine panel data that for some cohorts cover the entire careers we can directly model the changes in the hazard of job ending. We do this using a competing risks model and analyse separately exits to other jobs and exits to non-employment. We are primarily interested in the variation of these hazards over time but also account for the effects of elapsed duration, age and gender on the hazard of job ending.

2. Data

Our data are based on individual pension contribution records of workers covered by the Employee Pension Insurance Scheme. Data were originally collected for calculating pension accruals.

The Finnish pension system is a defined-benefit system where each employment spell contributes to the old-age pension with a fixed accrual rate multiplied by the duration of the employment spell. Current system was created in 1962 and reformed in 2005. Pension contributions are mandatory and employers are liable for arranging pension insurance for all their employees. The system is operated by private insurance companies. When the workers retire the full pension is paid by the insurance company that had insured the last employment spell. The Finnish Centre for Pensions supervises the system and acts as a clearinghouse that allocates pension liabilities to the companies that received the pension contributions and transfers the funds to the company that will pay the pension to the retiree.

To perform its task the Finnish Centre for Pensions has access to full pension contribution records of all pension insurance companies. Importantly for this study, these records include information on the starting and ending dates of all employment spells. The data set that is used for this study was created at the Finnish Centre for Pensions for developing indices that are used in converting the pension accruals to the price level prevailing at the time when the employees retire.

The sampling frame includes all individuals who have contributed to the pension system between 1963 and 2004. The sampling was done using a stratified design first selecting those born on the eighth day of each month from every fifth cohort born between 1905 and 1975. Within each cohort individuals were then picked at random until a desired sample size was reached. For the individuals that are selected into the sample all employment spells between 1963 and 2004 are included in the data. Data contain cohort-specific sampling weights that are used throughout the paper.

In the Finnish pension system employment spells contributed to the pensions from age 23 onwards. This age limit was reduced to 18 in 2005. For this reason the data from the earlier years only contain employment spells from age 23 onwards. For the last three cohorts employment spells are recorded from age 18 onwards. However, data contain the starting date of the employment spell also for the spells that were ongoing on the 23rd, and for the last cohorts, on the 18th birthday.

The first observations in the data are from 1963. Also in this case the data contain starting dates of all employment spells that were ongoing on January 1st 1963 no matter when the job had started. Therefore, the sample is representative of all ongoing employment spells between 1963 and 2004 for the cohorts that are included in the sample. Naturally the follow-up period varies between cohorts. The oldest cohorts are only observed at the end of their careers and the youngest cohorts only at the beginning of their careers. The 1940 birth cohort is the only cohort that is observed from age 23 to age 64. The age range when each cohort is observed and the sample sizes available for each cohort are displayed in Table 1.

Table 1 The observation period for each cohort in the original data set

Birth cohort	Years when observed	Ages when observed	Sample size	
			Individuals	Job spells
1905	1963–1970	58–64	283	324
1910	1963–1975	53–64	295	422
1915	1963–1980	48–64	296	510
1920	1963–1985	43–64	331	658
1925	1963–1990	38–64	401	1,082
1930	1963–1995	33–64	442	1,593
1935	1963–2000	28–64	469	2,863
1940	1963–2004	23–64	456	2,878
1945	1968–2004	23–59	494	4,569
1950	1973–2004	23–54	514	4,955
1955	1978–2004	23–49	538	4,563
1960	1983–2004	23–44	533	5,890
1965	1983–2004	18–39	1,022	13,094
1970	1988–2004	18–33	1,020	13,009
1975	1993–2004	18–29	975	10,223

The main limitations of the data are due to changes in the pension coverage over time. First, as already noted, the cohorts born before 1965 are only included in the data from their 23rd birthday. To keep the data consistent across cohorts, we have excluded all job spells ending before age 23 also for the younger cohorts. Second, the data excludes very short spells lasting for less than one month because these jobs were not insured under the Employees' Pension Scheme (TEL). Before 1965 this limit was six months and between 1965 and 1971 four months. We have no data from jobs insured under Temporary Employees' Pension Act (LEL) that cover the employees in construction, agriculture, forestry and harbour work. The main reason for this is that LEL-insurance is based on monthly gross earnings and has no need to record employment dates.

Finally, the data covers both public and private sector workers but the coverage of the public sector is incomplete before the 1980s. To minimize the effect of changes in insurance coverage, we exclude all public sector employment spells from most empirical calculations. In the empirical analysis we also exclude self-employed and farmers (both covered by their own pension schemes) and hence focus on the private sector employees covered by Employees Pension Scheme (TEL) for which we have information reported in a consistent way for the whole 42-year-period. Our analysis sample covers most private sector workers. For example, in 2004 about 90 per cent of private sector employees paid contributions to the Employees Pension Scheme.

In addition to dates of job spells the data contain only a limited amount of other information. Age and gender can be inferred from the id-codes. Reasons why job spells ended can be used to identify those who retire, but not to distinguish between dismissals and quits. As a partial solution we can classify jobs endings as quits and layoffs based on whether the employee started a new job within two weeks after the end of the previous job.

Other than removing the public sector employees, we have made only minimal adjustments to the original data. We removed short overlapping job spells in cases when a short spell begins and ends while a longer spell is ongoing. Job spells that are ongoing on December 31st, 2004 or ongoing on the day when the worker turns 65 are marked as censored. Jobs that were ongoing in January 1963 are also included in data and coded according to their original starting date. Data are left-truncated since we can only observe spells that lasted until 1963. Left-truncation also arises because spells that end before the 23rd birthday as well as spells lasting for less than 6 months before 1965, less than 4 months before 1972 and less than one month thereafter are not observed.

3. Comparison to other data sources

Most commonly studies on job stability are based on consecutive cross-sectional surveys that have been widely available for research purposes. Several studies in the United States have used data from the Current Population Survey (CPS). It collects information on tenure with the current employer or in the current job in various mobility supplements conducted at irregular intervals since 1951.

Comparable cross-sectional survey data have also been used in several other countries to measure the changes in job stability. Heisz (1999, 2005) used monthly tenure data collected in the Canadian Labor Force Survey (CLFS) since 1976, Gregg and Wadsworth (1995, 2002) and Burgess and Rees (1996, 1997, 1998) used data from the British Labor Force Survey (BLFS) and data from the General Household Survey (GHS). Both data sets include annual information on the current tenure of the respondents since 1975 and 1974, respectively. Vejsiu (2001) used data from the Swedish Level of Living Surveys and the Swedish Labor Force Survey. These two data sources contain information on the current tenure in ongoing jobs in Sweden from 1968 onwards. The Swiss Labor Force Survey used by Sousa-Poza (2004) has gathered information on the current tenure of the respondents since 1991.

While independent cross-sectional surveys are designed to be representative for the target population in each cross-section, the fact that they lack information on eventual tenure after the interview date is unfortunate. CPS income supplements offer a slight improvement making it possible to follow respondents in two consecutive years (Stewart 2002). Similar two year panel has been available in the aforementioned Swiss Labor Force Survey. Nevertheless, the main approach employed in studies using CPS data and comparable sources from other countries has been to analyze the changes in the current tenure distribution (Farber 1995, 1997, 2007; Jaeger & Stevens 2000) or to rely on strong assumptions that allow predicting future retention rates (Swinnerton & Wial 1995; Diepod, Neumark & Polsky 1997; Neumark, Polsky & Hansen 2000;) or eventual competed duration of the job held by the respondents (Hall 1982; Ureta 1992).

Several studies have employed panel data to analyze the changes in job stability. In the United States, Marcotte (1999), Polsky (1999), Gottschalk and Moffit (2000) and Jaeger and Stevens (2000) all use the Panel Study of Income Dynamics (PSID). Since 1976, the PSID has included a question on the current tenure, and in principle the annual interviews allow one to track employment histories of the survey respondents.

Other panel data sets used in US studies on job stability include the Survey of Income and Program Participation (SIPP) and the National Longitudinal Surveys (NLS). The SIPP data used by Gottschalk and Moffit (2000) and Bansak and

Raphael (2006) include information on the job histories of the individuals typically over a period lasting 32 months. The National Longitudinal Survey of Young Men and the National Longitudinal Survey of Youth, have been used for analyzing job stability by Bernhardt, Morris, Handcock and Scott (2000). These two data sets provide the researchers with a 16-year follow-up of the employment histories of young men starting from a cohort born in 1944, and include employer coding that captures job changes.

The German Socio-Economic Panel (GSOEP) was used by Winkelmann and Zimmermann (1998) and Bergemann and Mertens (2004). These data are available since 1984 and include information on both the current tenure of the individuals and their labor market transitions between consecutive interviews. In addition, when the GSOEP was initiated in 1984, the respondents were asked about the number of employers they had during the past 10 years. Even longer retrospective data is available in Britain. Booth, Francesconi and Garcia-Serrano (1999) use retrospective employment history data gathered in 1993 as a part of the British Household Panel Survey (BHPS). This survey asked the respondents to list their employment history until September 1990 starting from the day they left full-time education.

Naturally, the risk in using retrospective data for analyzing changes in job stability is that the respondents may not recall short employment spells in distant past which could lead to a false impression of declining job stability. A perhaps better alternative is to compare retrospective information in surveys conducted at different times. This approach was applied by Stevens (2005) who used data from the Retirement History Survey (RHS), the National Longitudinal Survey of Older Men (NLSOM) and the Health and Retirement Survey (HRS) to measure changes in the distribution of the longest job held by the respondents during their career.

Data from administrative registers has been used only recently for analyzing changes in job stability. For example, Bratberg, Salvanes and Vaage (2006) use linked employer-employee data from the period 1986-2002 to investigate the changes in job stability in Norway. Also Mahringer (2005a, 2005b) uses administrative employment spell data based on the Austrian social security records from 1975 onwards. These data sets, along with the pension register data employed in this paper, can be seen as considerable improvements to the job stability literature by making it relatively easy to follow job spells over time and by avoiding several problems researchers have previously faced with survey-based data.

4. Descriptive analysis

Changes in average tenure over time

As a first attempt to describe changes in job stability in our pension records data we examine the changes in the average tenure over time. We pick an arbitrary date, October 15th, each year and report the average elapsed duration of the jobs that were ongoing on that day. The results are plotted in Figure 1. We report both simple averages and regression adjusted averages that account for changes in the age and gender distribution. The horizontal line in the figure refers to the overall arithmetic average, 8.7 years.

The unadjusted numbers are trending slightly upwards over time. This partly reflects population aging. Older workers have accumulated on average longer tenure, and an increase in their share in the data increases the average tenure. It is not clear whether any adjustments for population aging should be made; an increase in average tenure is a real phenomenon even if it is caused by a change in the age structure. However, our sample is not exactly representative of the current population in any single year; it is only representative for the certain cohorts. Since a new cohort is added to data every fifth year, the sample gets successively older for five years and then suddenly younger as a new cohort enters in the data. The effect of sampling scheme can be seen as a modest five-year cycle in the unadjusted average tenure. The age effect is particularly strong in the end of the data. The youngest cohort born in 1975 is included in the data in 1998, but after that no new cohorts enter and the sample gets gradually older.

The largest changes in the adjusted average tenure series occur in the end of 1970s when the average tenure increases and in the 1990s when average tenure starts a gradual decline.

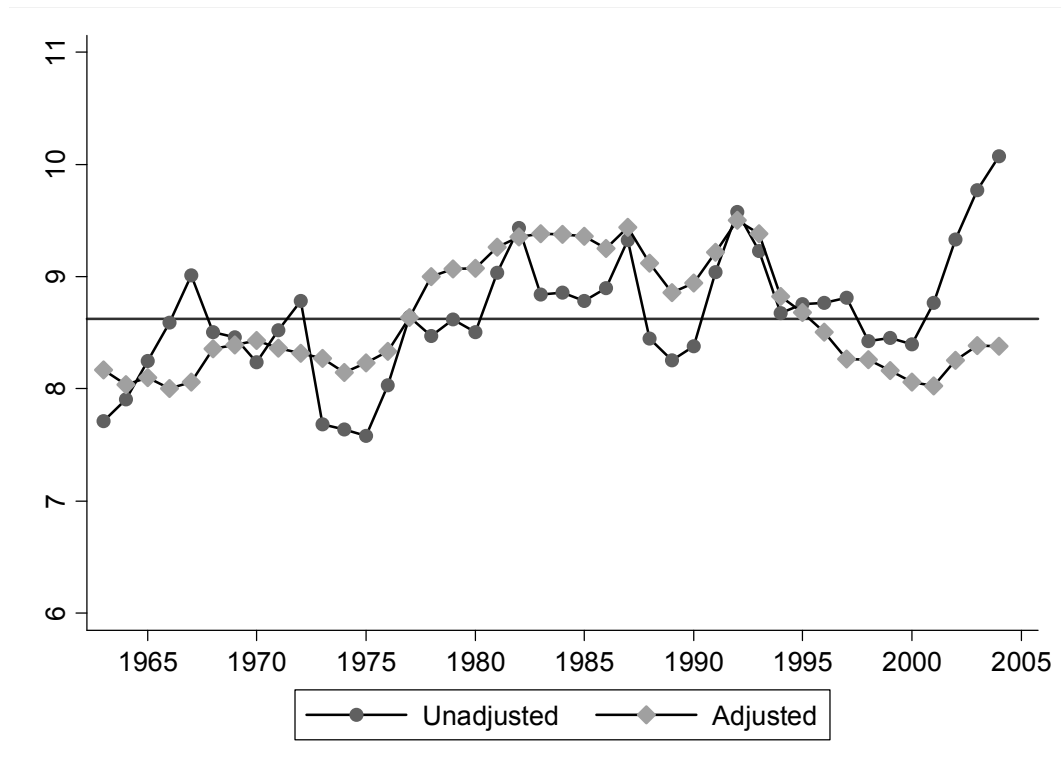


Figure 1 Average tenure in current job 1963–2004

Note: Adjusted series is created by regressing elapsed duration on gender and one-year age dummies and calculating the predicted values when gender and age distributions are set equal to the sample average.

Comparison of register vs. survey data

To verify that changes occurring in the register data are not due to changes in reporting procedures or changes in insurance coverage, we compared the tenure distribution in the register data to the figures calculated from the Labor Force Survey¹.

Comprehensive micro-level data from the Labor Force Survey are not available for research purposes in Finland. Therefore, we cannot calculate simple statistics such as the mean tenure for years before 1997. However, Statistics Finland regularly publishes monthly data on new jobs, defined as jobs with tenure less than a year. Statistics Finland also provided us with unpublished tabulations on the tenure distribution that allow calculating the fraction of jobs with current tenure of more than ten years in a consistent way.

¹ Questions on the current tenure were first added to the annual interview of the Labor Force Survey in the fall of 1982. The question on current tenure was included in the survey every year until 1987 and then and every other year between and 1987 and 1993. In 1995 and 1996, the tenure question was included in the EU Labor Force Survey conducted in spring. From 1997 onwards the question on elapsed tenure has been included in the monthly Labor Force Survey. In addition to changes in survey dates, also the survey question has been changed slightly which may make the LFS time-series less consistent over time.

In Figure 2 we plot the fraction of workers with elapsed tenure over one year and in Figure 3 fraction of workers with elapsed tenure over ten years based on both the pension records and on the survey data. We have made the sources as comparable as possible. Since the annual interview of the LFS was held between September and December in the 1980s we use data from the last quarter of the year also after 1997. After 1980s the coverage of the register data is better so that, in contrast to the other analyses in this paper, also the public sector employees and workers under 23 can be included in this comparison.

According to Figures 2 and 3 the evolution of the tenure distribution is very similar in the survey and register data. According to both sources, on average, about 80% of the workers have been in their jobs for more than a year. Also the changes in tenure distribution seem similar. According to both sources there was a large increase in the fraction of workers with more than one year of tenure in the beginning of 1990s. This does not imply that job markets were more stable during those years, rather the opposite. Finland experienced a major recession in the beginning of 1990s and very few workers were recruited during those years. Hence the fraction of new workers with short tenure declined and average tenure increased.

Survey and register data seem to produce similar numbers also in Figure 3 that displays the fraction of workers with more than ten years of tenure. According to either source about 35 to 40% of workers have been working for their current employer for more than ten years. There is also a slight increase over time, again potentially explained by population aging.

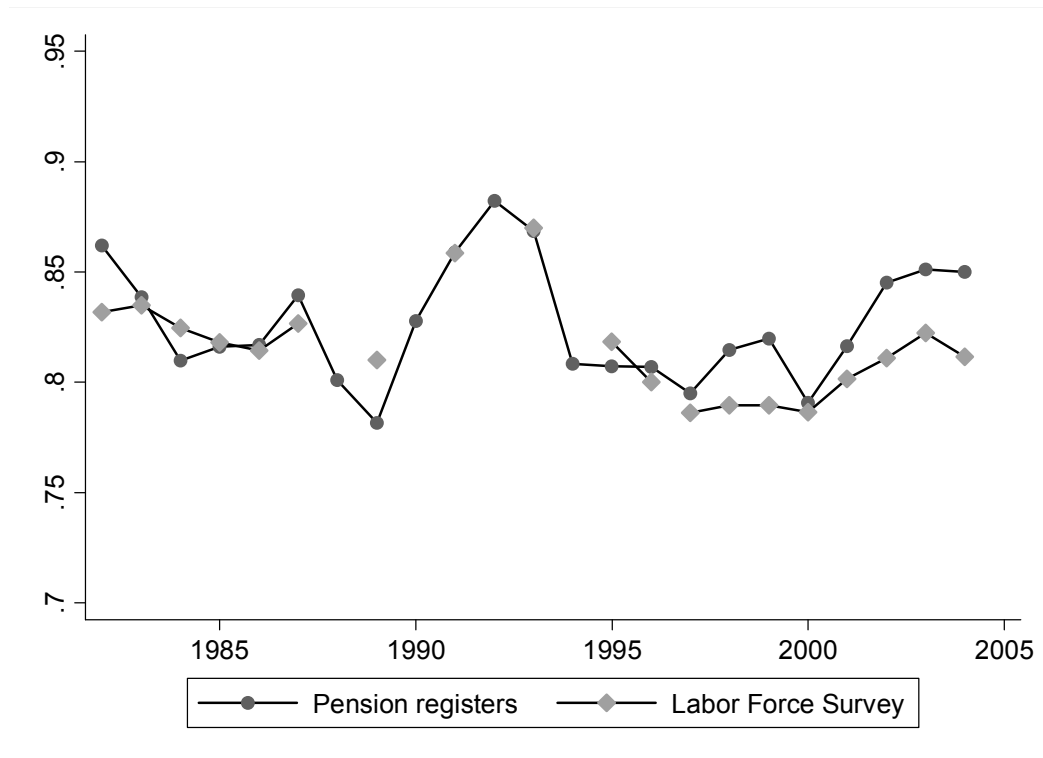


Figure 2 Share of workers with elapsed tenure > 1 year in register vs. survey data

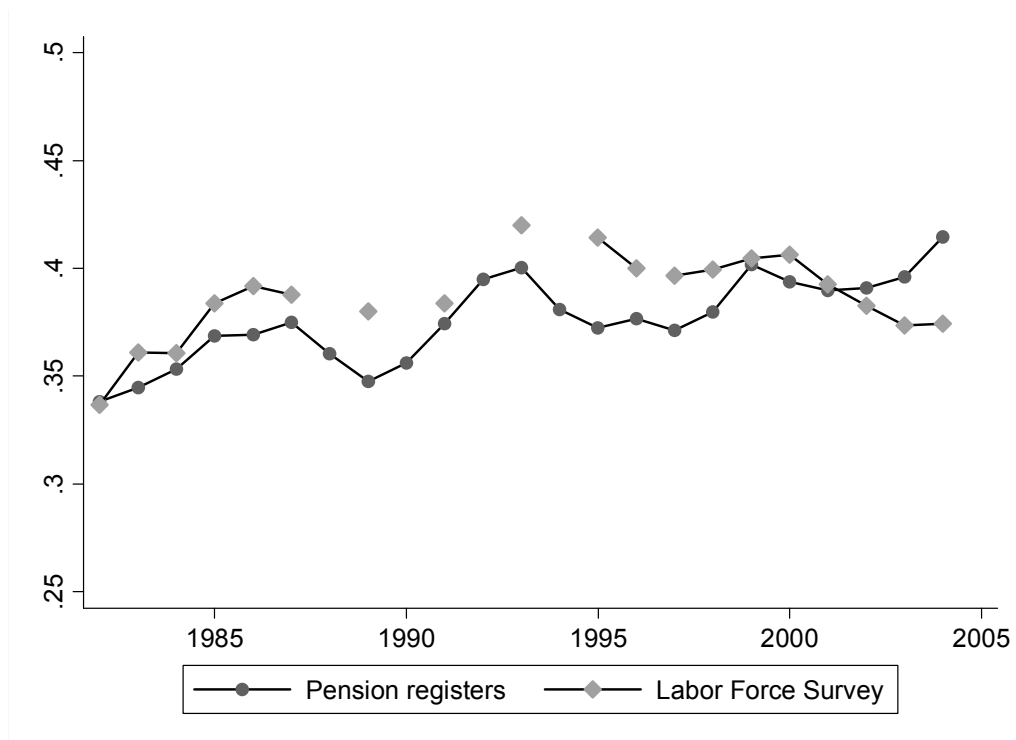


Figure 3 Share of workers with elapsed tenure > 10 years in register vs. survey data

Notes to Figures 2 and 3:

Elapsed tenure in register data is calculated based on jobs ongoing on October 15th each year. Data includes workers between 18 and 64 and cover both public and private sector employees. Self-employed are excluded. Survey data up to 1993 is based on annual interview of the Labor Force Survey. Data from 1995 and 1996 are from EU Labor Force Survey. From 1997 onwards data is based on monthly Labor Force Survey. To avoid inconsistencies due to differences in survey dates we have used the numbers from the last quarter of each year.

Differences across cohorts

Examining changes in current tenure in a given year may not be the best way to capture long-term changes in job stability. Possibly a better approach is to compare differences in elapsed tenure at a given age between consecutive birth cohorts.

In Figures 4a and 4b we follow the example by Farber (2007) and plot average elapsed tenure by age for different cohorts separately for men and women. For clarity we aggregate the cohort data slightly and take an average of three consecutive cohorts so that the first line refers to cohorts born between 1905 and 1915 and the last line to cohorts born between 1965 and 1975.

According to Figures 4a and 4b the average elapsed tenure increases almost linearly by age. The increase is more rapid for men. At age 50 the average tenure for men is about 15 years and average tenure for females about 12 years. Differences across cohorts are not very large except in the older ages where also the number of observations gets small and the estimates less precise. Still, one can note that the youngest cohorts born between 1965 and 1975 have accumulated slightly less tenure at any given age than the cohorts born before 1965.

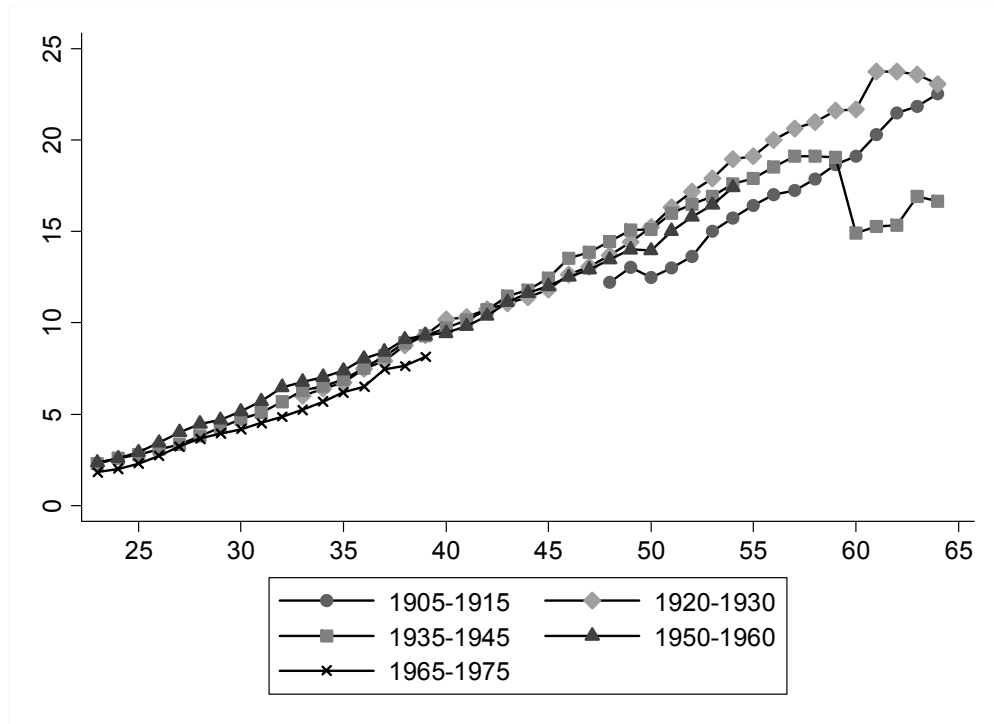


Figure 4a Mean elapsed tenure by age and cohort, men

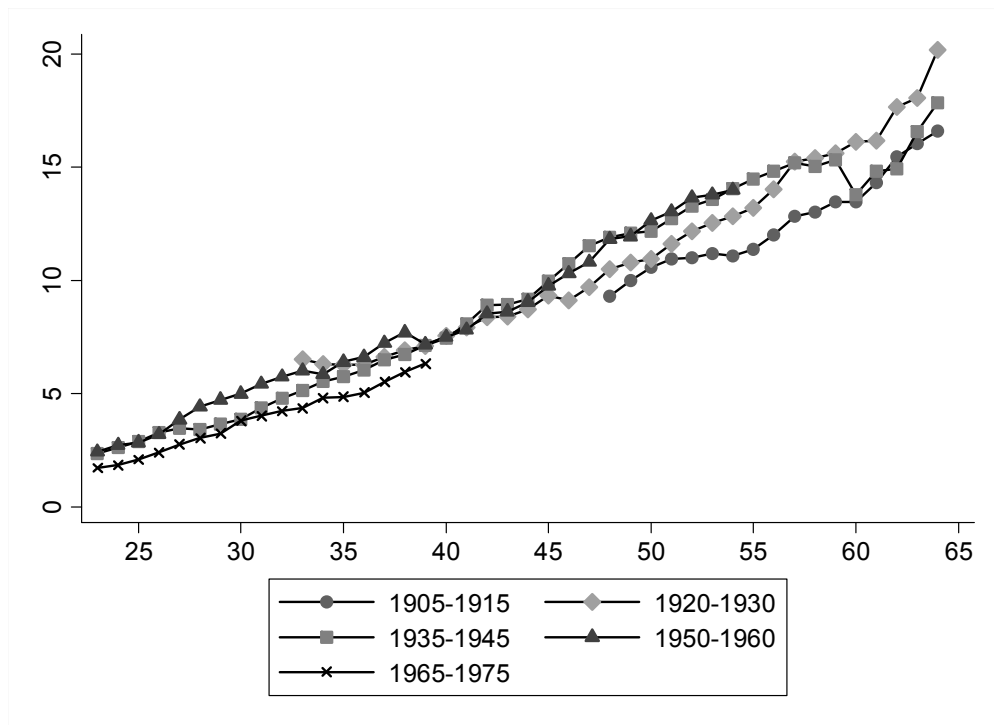


Figure 4b Mean elapsed tenure by age and cohort, women

The differences across cohorts can be seen more clearly from Figure 5 that presents the cohort effects from a regression model where (log) current tenure is explained by a set of cohort dummies and a full set of one-year age dummies.

The figure shows that, conditional on age, average tenure was lower in the first female cohorts but then remained rather stable for both men and women until the cohorts born in 1950s. For both men and women age adjusted average elapsed tenure is smallest among the youngest cohorts.



Figure 5 Cohort effects on current tenure

It should be noted that the decrease in the average age-adjusted tenure does not necessarily imply that the jobs held by the youngest cohorts would be less stable but also reflects a delay in the age of entry into the labor market due to an increase in time spent in education. Figure 6 displays the labor force participation rate and the employment rate in age group between 15 and 24. Both fluctuate according to the business cycle but have also declined substantially over time. The labor force participation rate in this age group is about ten percentage points lower and the employment rate about twenty percentage points lower in 2004 than in 1970.

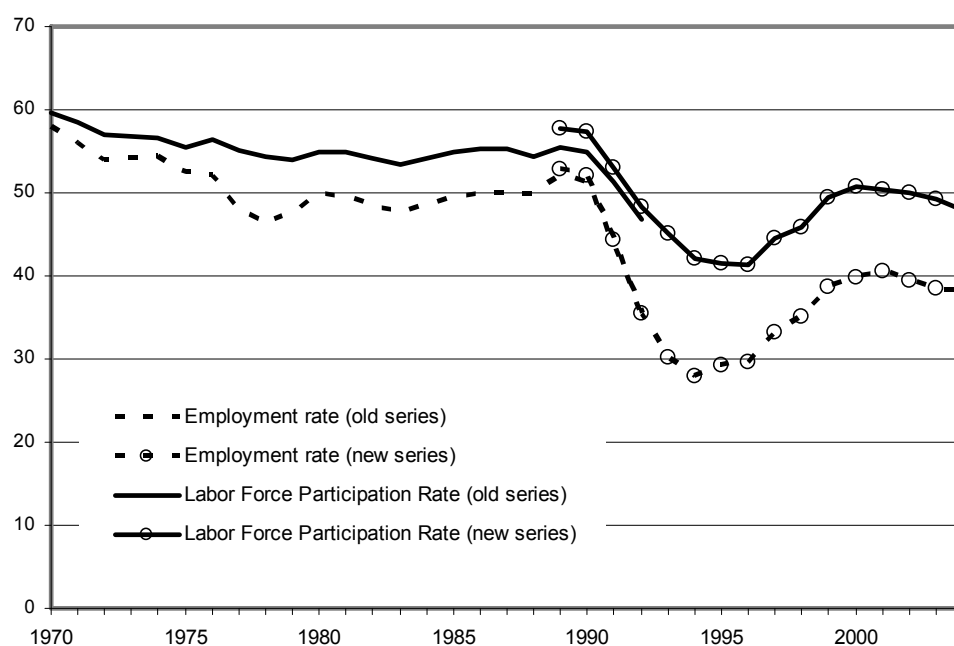


Figure 6 Labor force participation rate and employment rate in ages 15–24

Data source: Labor Force Survey, Statistics Finland database. Break in series is due to the reform in the Survey.

5. Effects of the changes in entry rates and the risk of job loss on average tenure

Though commonly used, the change in mean elapsed duration of ongoing employment spells may not be a particularly informative statistic for measuring the *changes* in job stability. To illustrate this we simulate the dynamic effects of changes in job stability defined as the change in the risk of job ending on average tenure.

We follow Lancaster (1990) and start by noting that the number of persons employed at time t_0 is

$$\int_0^{\infty} n(-y)S_{-y}(y)dy, \quad (1)$$

which depends on the number of entrants y periods ago $n(-y)$ and their historical y -period survival rates $S_{-y}(y)$. Hence, the mean elapsed duration of ongoing spells at time t_0 is

$$\mu_t = \frac{\int_0^{\infty} n(-y)S_{-y}(y)y dy}{\int_0^{\infty} n(-y)S_{-y}(y)dy}. \quad (2)$$

In general, the expression (2) depends on all past entry rates and all past survival rates. For example Ureta (1992) notes that calculating average completed tenure based on survival rates calculated from a cross-section data as was done in the famous paper by Hall (1982) leads to a bias if the arrival rates are not constant. Ureta's example of non-constant arrival rates had to do with increased labor force participation rates by women. Similar effects could be caused by large scale changes in immigration, changes in school-leaving age or major swings in the business cycle.

However, even if the entry rates were constant, the mean elapsed duration of ongoing jobs depends on all past survival rates and not only on the recent changes that a measure of changes in job stability should capture. Hence, for example a major recession that causes a temporary shock to the job exit rates affects average tenure long after the recession has ended.

In Figure 7a we illustrate the effect of an increase in the labor market entry rate on the average duration of ongoing jobs. We start from a stationary state where the number of new entrants is constant and equals the labor market exit rate. We estimate a Kaplan-Meier survival function from our pension record data and use

expression 5 to calculate average duration of ongoing jobs. To simplify calculations, we assume that there is no unemployment so that the workers whose contracts end immediately find a new job where their tenure is naturally initially zero.

We then increase the entry rate with a constant number of new entrants each year so that total employment increases by 20 per cent in 20 years and remains constant thereafter. We use equation (2) to calculate average tenure in ongoing jobs each year from $t = 0$ to $t = 100$. According to the results presented in Figure 7a the increase in labor market entry decreases the average tenure because there are more recent entrants with short tenure. Interestingly the effect of an increase in the entry rate has a long-lasting effect. Even at time $t = 30$ the average tenure is substantially below the initial level even though the increase in the entry rate had ceased ten years ago. The average tenure stabilizes to the initial level after about forty years.

A researcher comparing average tenure generated by our simulation between, for example, years $t = 0$ and $t = 30$ might well conclude that average tenure has declined and jobs become less stable. The conclusion would be quite misleading. The survival function of new jobs, and hence the average eventual completed tenure, is constant in our simulations. The decrease in average elapsed tenure in a cross-section is entirely due to an increase in the number of recent entrants.

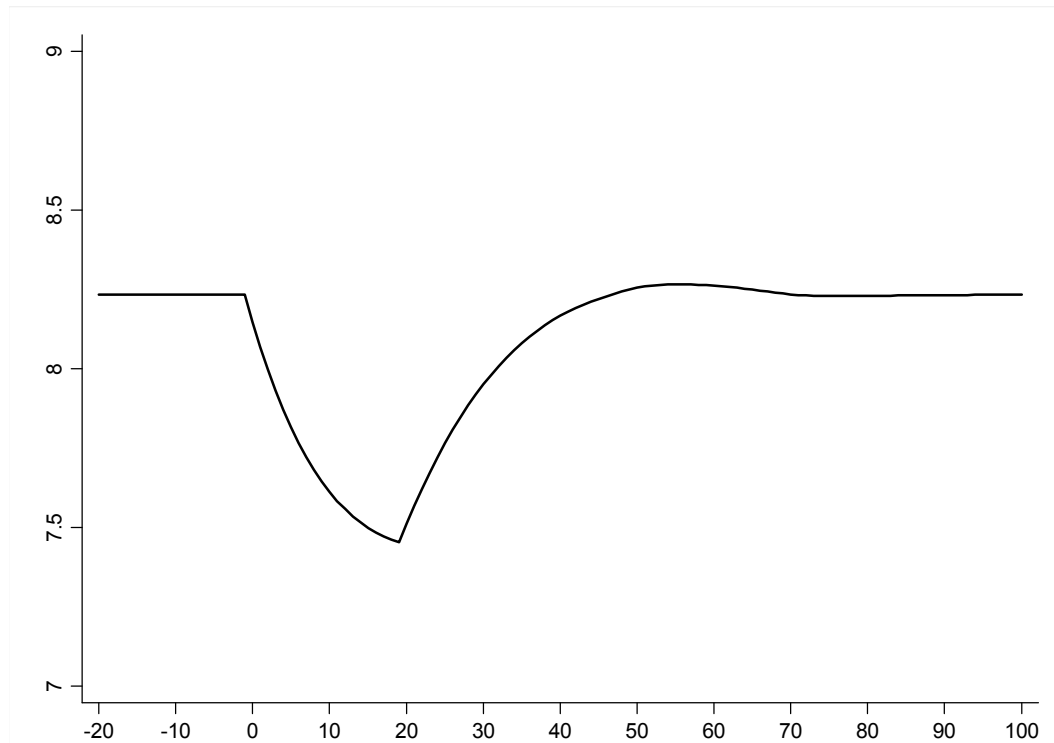


Figure 7a *Simulation results: Effect of an increase in entry rate on average tenure*

In Figure 7b we present results from another simulation more closely related to changes in job stability. We keep the number of new entrants constant but simulate the effects of a temporary increase in the hazard of job ending. We start again with survival function estimated from our pension records data but increase the hazard of job loss by 50 percent for five years for all cohorts that are in the labor market between years $t = 0$ and $t = 5$. After five years we return the hazard rate to its initial level. Average elapsed tenure is again calculated based on equation (2) for time periods from $t = 0$ to $t = 100$.

As shown in Figure 7b the increase in the hazard of job ending decreases the average tenure. Note that this is partly due to our simplifying assumption according to which the workers who lose their jobs are immediately re-employed (with zero tenure). This would also imply that hiring rate would have to increase. If an increase in the job ending rate led to an increase in unemployment, the change in average tenure would depend on the changes in re-employment rates. In a typical recession the re-employment rate decreases and leads to an increase in average tenure (Burgess and Rees 1996). The number of new (low tenure) jobs typically starts to increase only when recession is over. It is therefore better to interpret our simulation as an impact of temporarily higher volatility rather than as an effect of recession.

The most interesting feature of simulation results presented in Figure 7b is that a temporary increase in volatility has long-lasting effects on average tenure. Even though the shock to the job ending rates only lasts for five years the average tenure is substantially below the initial level several decades afterwards. Again data on the mean elapsed duration of ongoing jobs in consecutive cross-sections would give a misleading picture on the changes in uncertainty.

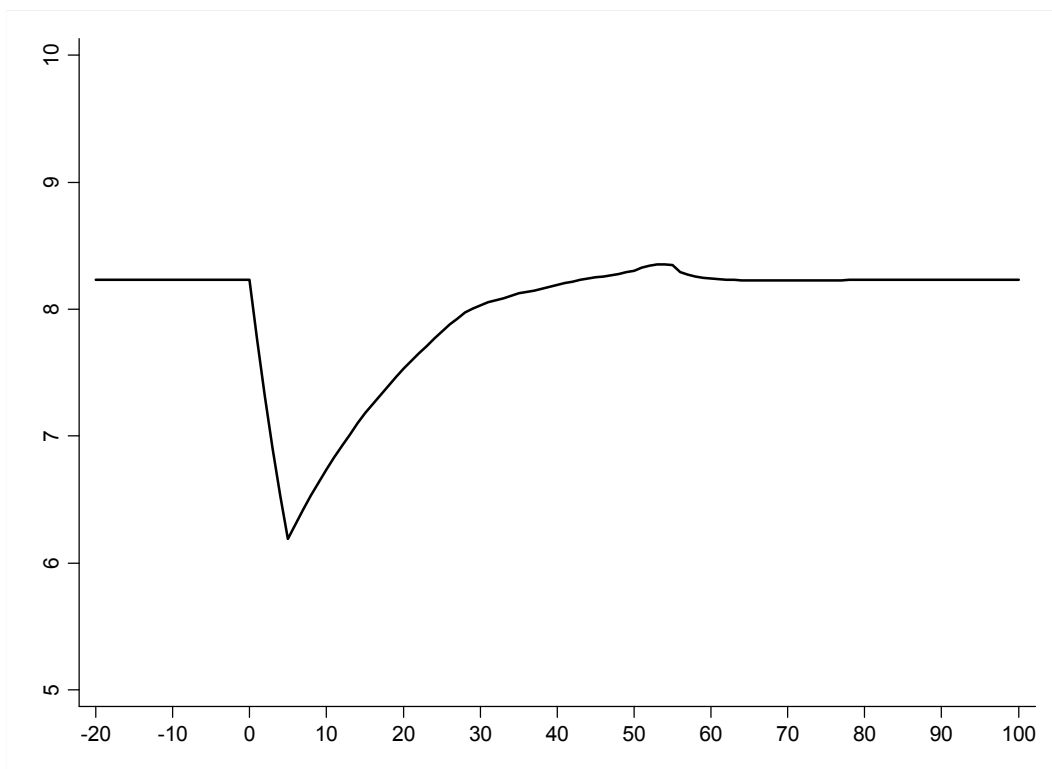


Figure 7b *Simulation results: Effect of an increase in the hazard of job ending on average tenure*

6. Changes in the risk of job loss – Results from a duration model

The most natural measure of labor market uncertainty is probably the risk that a job ends. A standard way of modeling this risk is the conditional probability that a job ends given that it has lasted for a given time, i.e. the hazard rate. Modeling the effects of exogenous covariates on the hazard rate - instead of on e.g. the completed duration of a job - also provides a simple ways to account for censoring due to a finite observation period, duration dependence, and the effects of covariates that vary over time.

Our data is a sample of jobs ongoing on January 1st 1963 and of jobs beginning between 1963 and 2004. It is therefore a mixture of stock and flow samples. Jobs that began before 1963 are observed only if they are still ongoing in 1963 creating a left-truncated sample. Left truncation arises also because we observe only employment spells ongoing after the 23rd birthday, and because jobs lasting less than 6 months until 1965, 4 months until 1972 and a month thereafter are not included in the data. The starting date is known also for the stock-sampled jobs. The ending dates of jobs that are still ongoing in the end of 2004 are unknown, and the data is hence right-censored at this point. We also censor jobs on the 65th birthday as workers over 65 are no longer paying pension contributions and are thus not followed in the data.

To model the changes in the hazard of job ending we use a competing risks model with two possible destinations: non-employment and a new job. Our definition is based on whether we observe a new employment spell within two weeks after the previous spell ends. We also experimented with a time limit of four weeks for this definition, but this had practically no effect on the results.

A large number of jobs end soon after they begin. After a job has lasted for a year, the hazard of job loss declines to about a fifth of the hazard prevailing during the first four months. The decline in the hazard of job change decreases also though the decline is smaller. To allow flexible forms of duration dependence in both hazards we specify a piecewise constant baseline hazard function where the hazard stays constant for four month periods during the first year and for one year periods thereafter.

We explain the changes in the hazard rates by gender, age and time. Again we aim to maximum flexibility and include the time varying covariates as a set of one-year age and time dummies. We assume that the covariates have proportional effects on the hazard rates. We also assume that there are no interactions so that, for example, duration dependence is independent of age. While these assumptions might be questioned, we would argue that they have little impact on

our primary objective; consistent estimation of the time effects that capture the changes in the hazard of job ending over time.

We account for individual-specific unobserved heterogeneity by specifying a mixed proportional hazard model with three discrete points of support following the approach by Heckman & Singer (1984). The choice of the number of support points is somewhat arbitrary, but experimentation with different numbers of discrete points revealed that the other parameter estimates remained practically unchanged with two, three, or four mass points. However, imposing a restriction that there is no unobserved heterogeneity would have substantial effects on the estimates.

We use the partial likelihood method (Lancaster 1979) that makes accounting for left truncation relatively easy. Jobs that are ongoing at the start of the observation period contribute to the hazard estimates only from the entry date onwards. The jobs that end within the observation period contribute to the likelihood function through both the hazard and the survival function, but the jobs that are still ongoing at the end of the observation period contribute only through the survival function.

Our model specification can be formalized as follows. We define the hazard function related to destination s for job j of individual i conditional on the vectors of observed covariates x_{ij} and individual-specific unobserved heterogeneity terms $\nu_i = (\nu_{i1}, \nu_{i2})$ as

$$h_s(t_{ij} | x_{ij}, \nu_i) = \lambda_s(t_{ij}) \exp(x'_{ij} \beta_s + \nu_{is}) \quad (3)$$

We specify the baseline hazard function $\lambda_s(t_{ij})$ for destination s as

$$\lambda_s(t_{ij}) = \sum_{k=1}^K \lambda_{sk} I_{t_{ij} \in [d_{k-1}, d_k)} \quad (4)$$

where I is an indicator function splitting each job spell into K episodes.

Furthermore, the survival function related to destination s for job j of individual i given the vectors of observed covariates x_{ij} and individual-specific unobserved heterogeneity terms ν_i is defined as

$$S_s(t_{ij} | x_{ij}, \nu_i) = \exp \left[- \int_0^{t_{ij}} h_s(u | x_{ij}, \nu_i) du \right]. \quad (5)$$

Let θ denote the parameter vector to be estimated and w_i the sampling weight of individual i . In addition, let c_{ijs} denote a dummy variable indicating whether job j of individual i is censored. The pseudo log-likelihood function can now be written as

$$\log L(\theta) = \sum_{i=1}^N w_i \log \left[\sum_{m=1}^3 \pi_m \prod_{j=1}^{J_i} \prod_{s=1}^2 \frac{h_s(t_{ij} | x_{ij}, \nu_m)^{1-c_{ijs}} S_s(t_{ij} | x_{ij}, \nu_m)}{S_s(e_{ij} | x_{ij}, \nu_m)} \right], \quad (6)$$

where the term between the brackets is the marginal likelihood contribution of individual i . The parameters ν_m and π_m denote the mass point vectors and corresponding probabilities of the unobserved heterogeneity distribution that are estimated along with the other parameters of the model.

We parametrize the probabilities π_m of the unobserved heterogeneity distribution as

$$\pi_m = \frac{\exp(p_m)}{\sum_{m'=1}^M \exp(p_{m'})} \quad (7)$$

and normalize p_1 to zero. This parametrization takes automatically care of the requirement that the estimated probabilities must lie between zero and one.

The role of the division by $S_s(e_{ij} | x_{ij}, \nu_m)$ in the pseudo log-likelihood function is to take into account the left truncation in our data. That is, we condition the likelihood contribution of each job on the fact that it must have survived until e_{ij} which denotes the elapsed duration at which job j of individual i enters the data.

Our approach produces a large number of parameters. For expositional reasons we prefer presenting the estimated hazard ratios and their 95% confidence intervals in a graphical way. The parameter estimates and their standard errors are reported in the appendix. We start by plotting the baseline hazard function in Figure 8a. We omit the first category with elapsed duration of 0-4 moths. As shown in the figure the hazard of job loss declines rapidly during the first few years. After having lasted for seven years the hazard is about 10 per cent of the hazard during the first four months. After that the hazard of job loss remains approximately constant for twenty years. The hazard of job change declines also, but the decline is clearly slower. At very high durations the estimates of job change hazard get less precise due to a small number of job changes.

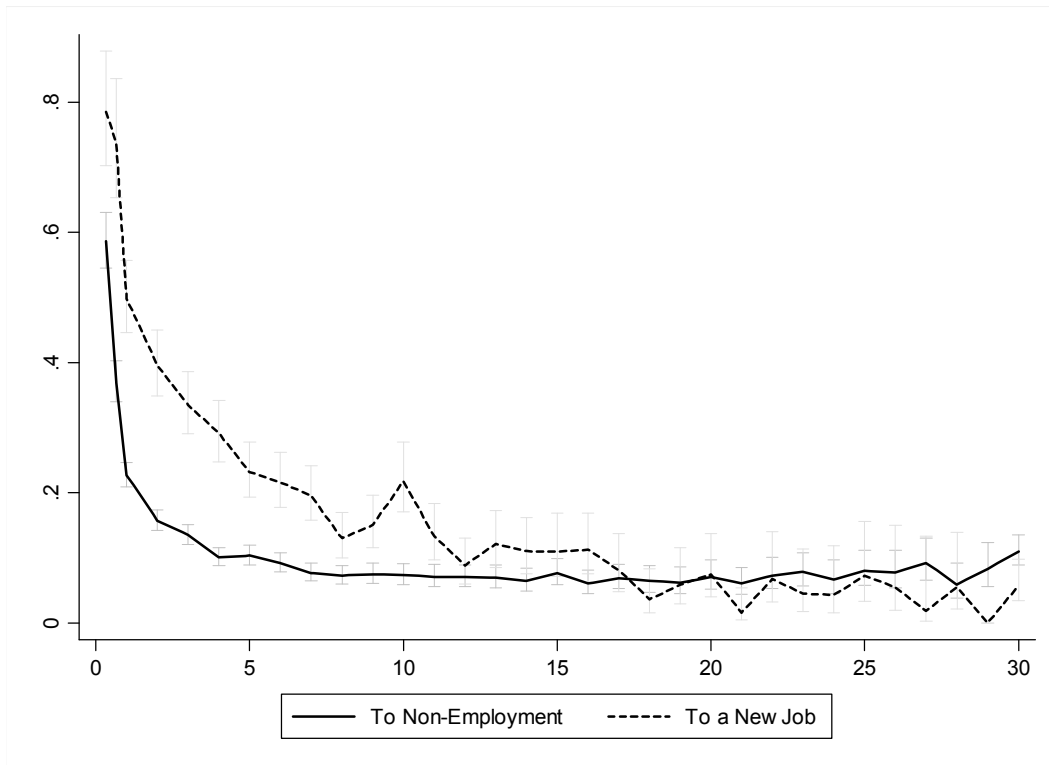


Figure 8a Hazard of job ending as a function of elapsed duration

Figure 8b presents the effect of age on the hazard rates. We have chosen age 23 as the reference category and omit ages 63 and 64 from the figure for expositional reasons. The risk of job loss first declines as the workers get older. From about age 52 onwards the hazard increases rapidly with age reflecting the effect of early retirement and gets very high after age 60. Hazard of job change declines almost linearly as the workers get older.

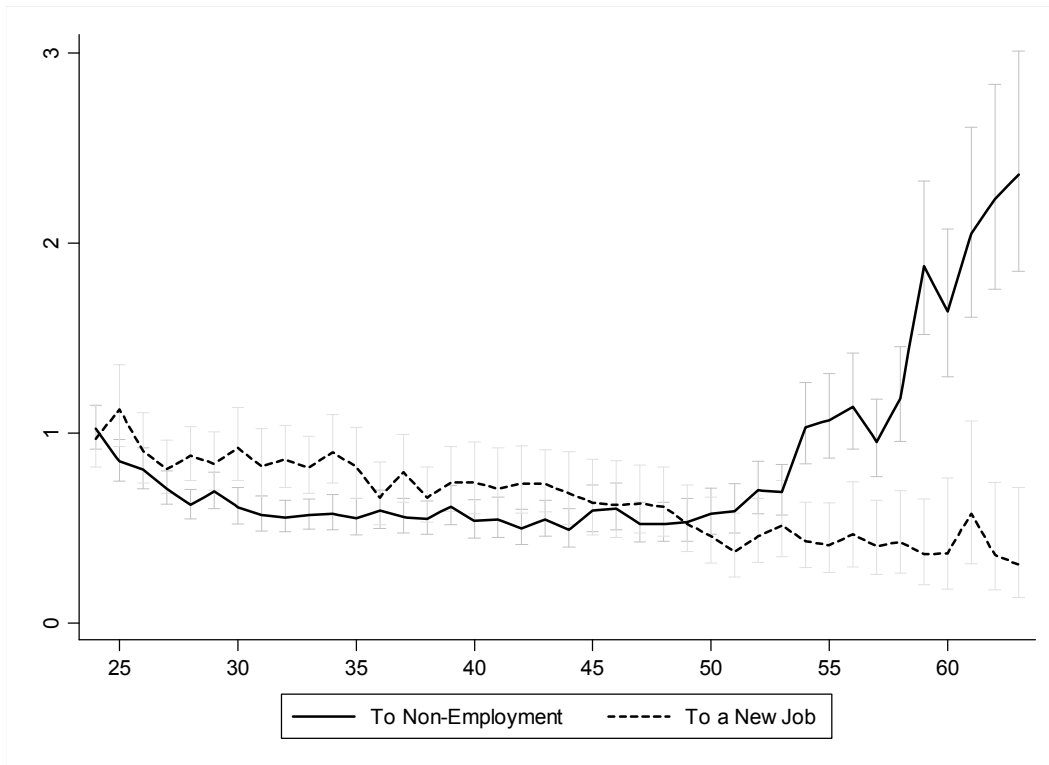


Figure 8b The effect of age on the hazard of job loss

Finally in Figure 8c we plot the time effects using 1963 as a reference period. We find that the hazard of job change is much more volatile than the hazard of job loss. Early 1970s, late 1980s and late 1990s were years of particularly rapid job-to-job mobility. In contrast, there is only one peak in the hazard to non-employment that coincides with the recession in the early 1990s. If we interpret the hazard to non-employment as an indicator of uncertainty, we can also note that the uncertainty after year 2000 is approximately at the same level as it was forty years earlier, in the beginning of the 1960s.

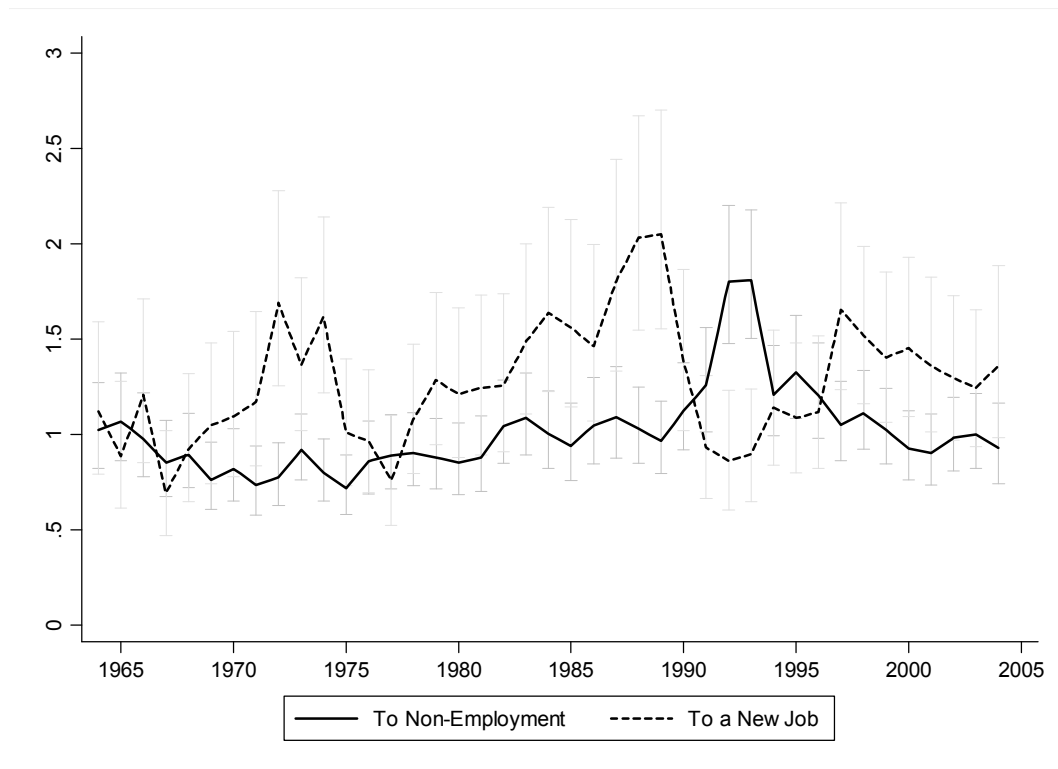


Figure 8c Time variation in the job termination hazard

A comparison of the job-to-job and job-to-non employment hazards to the mean elapsed duration of ongoing jobs presented in Figure 1 reveals that these series appear to be almost unrelated to each other. The increase in average tenure observed around 1980 could be partially explained by lower job-to-job mobility in the end of the 1970s and the decrease in average tenure after 1993 by an increase in job-to-job mobility in mid 1990s, but clearly something else is going on also. In addition, two possible measures of job security: mean elapsed duration of ongoing jobs and hazard of job ending point to very different time pattern in the changes of uncertainty prevailing in the labor market.

As we noted in the previous chapter, the mean elapsed duration of ongoing jobs depends not only on the current hazard of job ending, but also on past hazards and the variation in entry rates. To quantify these effects we used the estimates from a slightly simplified version of our duration model² together with the number of new entrants each year and the initial duration distribution that prevailed in 1963 to simulate the average tenure in each year.

As shown in Figure 9 the simulated average tenure correspond relatively well to the observed average tenure in any given year. Only major deviation between the observed and the simulated series occurs in the end of the series, and this can be

² The model used for the simulation exercise includes only duration dependence and year effects with no unobserved heterogeneity.

explained by the changes in the age distribution of our sample. To quantify the effects of changes in job ending hazards on average tenure, we then restricted both the job-to-job and job-to-non employment hazards at their sample averages and simulated again the implied mean elapsed tenure for each year. As shown in Figure 9 this has surprisingly little effect on the time pattern of the average tenure. However, if we also restrict the number of new entrants to a constant the pattern is very different irrespective of whether we keep the hazard rates constant or not. Note that even if both the entry rate and the hazard rates are constant, average elapsed duration generally changes over time since there is no reason to assume that the tenure distribution in 1963 reflected a steady state. In fact, the mean elapsed duration of ongoing jobs in 1963 was rather low, possibly due to earlier labor supply shocks.

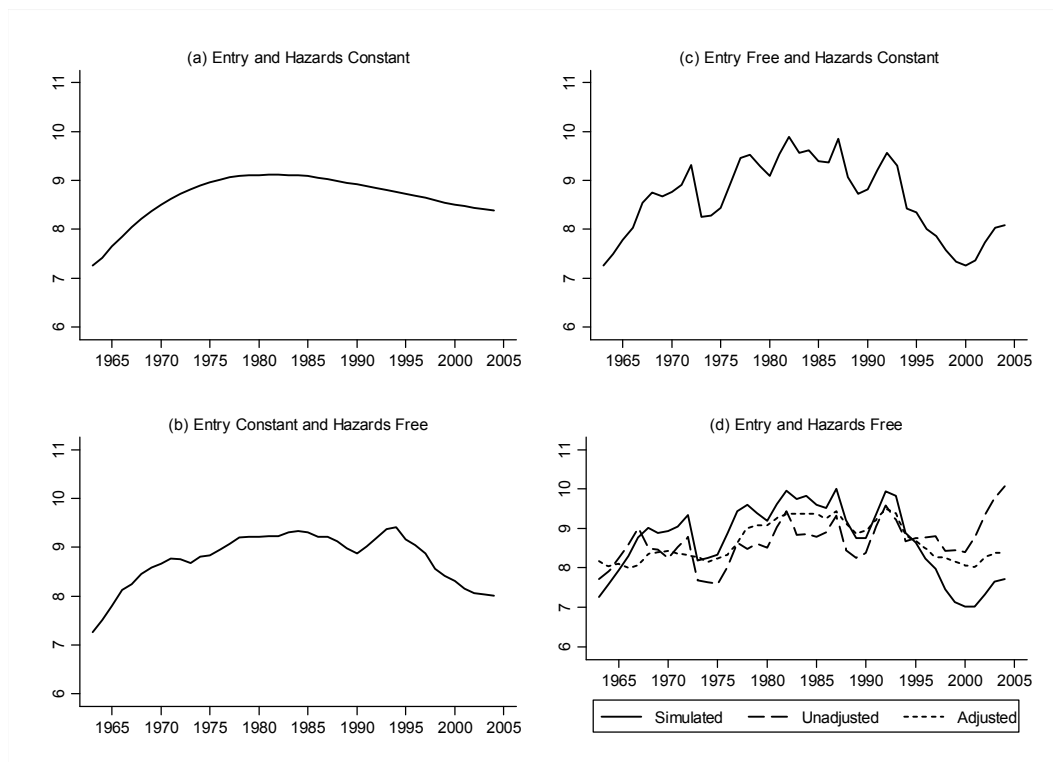


Figure 9 *Simulation results*

7. Lifetime jobs

For the cohorts that have retired by the year 2004 we can directly observe the duration distribution for the longest job over the career. In Figure 10 we present the average length of the longest job during the career for cohorts born in 1940 or before. For the cohorts for which we observe almost the entire career this can be done accurately. For the two oldest cohorts that are under observation from ages 53 or 58 onwards we can be less certain that the job we observe in the end of the career is the longest job the individual has held over lifetime. To make the data more consistent between cohorts we restrict the data to those who were employed at some point after age 50.

According to Figure 10 the average length of the longest job over lifetime in these cohorts was around 22 years for men and a few years shorter for women. For men the average length has been very similar across cohorts born between 1905 and 1940. For women there is a clear upward trend with the 1940 cohort reaching almost the same length as men. We have also calculated various quantiles of this distribution. For example the median durations are very close to the mean durations, both in terms of the level and changes.

Our results reveal that, at least for the cohorts that have already retired, long-term employment contracts were an important feature of the labor market. Also no indications that lifetime contracts had become less common can be seen in the data. Rather the long-term contracts seem to have gotten more common, at least for women, perhaps reflecting their increased participation into the labor market.

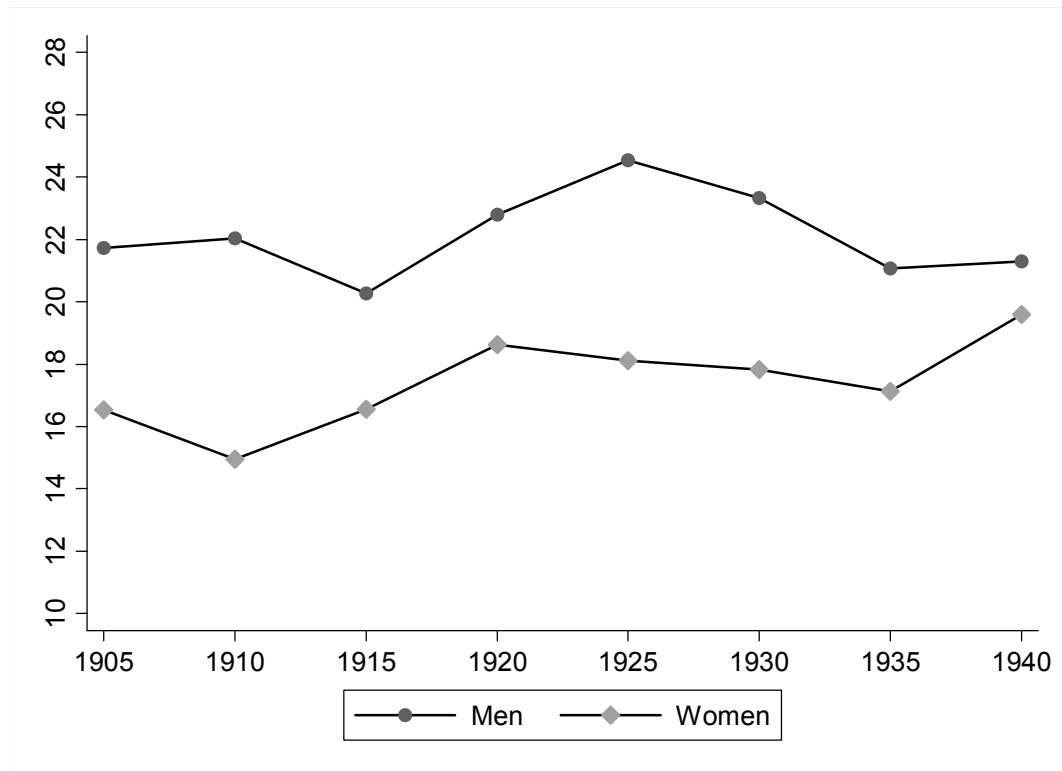


Figure 10 Average length of longest job during the career

8. Conclusions

The estimates presented in this paper suggest that there has been no long-term trend in the mean elapsed duration of ongoing jobs in Finland. Average tenure has still varied considerably over time, but this variation has been mainly due to the variation in the number of new entrants into the labor market. If we instead measure job market stability by the hazard of job loss - or more precisely by the hazard of job ending that leads into non-employment - using a standard duration model specification, we find that job stability declined during a major recession but that the risk of job loss after year 2000 is at the level observed forty years ago.

Our results suggest that there are important caveats to be kept in mind when using the elapsed duration of ongoing jobs to measure changes in job stability. The changes in average tenure appear to be mainly related to the changes in the number of new jobs rather than changes in stability of existing jobs.

This paper also illustrates the benefits of using register data originally collected for administrative purposes. Pension registers are particularly useful since calculating pensions typically requires information from the entire careers. Compared with previous research on job stability we can both cover much longer period starting from early 1960s to present and observe the entire lifetime job histories. The Finnish data is of unusual quality because of centrally coordinated pension system that requires information on dates of employment contracts. Still similar data probably exists with pension authorities also in other countries.

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Appendix

Parameter Estimates for the Mixed Proportional Hazard Competing Risks Model of Job Endings

	To Non-Employment			To a New Job		
	Estimate	SE	P-value	Estimate	SE	P-value
Constant	1.626	0.147	0.000	-0.938	0.203	0.000
Male	-0.195	0.033	0.000	-0.023	0.035	0.258
Duration dependence (reference category 0)						
4 months	-0.534	0.037	0.000	-0.242	0.057	0.000
8 months	-0.995	0.043	0.000	-0.303	0.063	0.000
1 year	-1.482	0.042	0.000	-0.697	0.057	0.000
2 years	-1.848	0.051	0.000	-0.927	0.065	0.000
3 years	-2.002	0.057	0.000	-1.094	0.073	0.000
4 years	-2.291	0.069	0.000	-1.234	0.082	0.000
5 years	-2.268	0.073	0.000	-1.462	0.092	0.000
6 years	-2.385	0.079	0.000	-1.532	0.099	0.000
7 years	-2.567	0.091	0.000	-1.631	0.108	0.000
8 years	-2.620	0.097	0.000	-2.038	0.135	0.000
9 years	-2.591	0.104	0.000	-1.892	0.136	0.000
10 years	-2.613	0.109	0.000	-1.525	0.125	0.000
11 years	-2.646	0.125	0.000	-2.012	0.162	0.000
12 years	-2.646	0.120	0.000	-2.424	0.199	0.000
13 years	-2.664	0.127	0.000	-2.108	0.179	0.000
14 years	-2.742	0.135	0.000	-2.203	0.195	0.000
15 years	-2.570	0.131	0.000	-2.208	0.219	0.000
16 years	-2.799	0.151	0.000	-2.182	0.205	0.000
17 years	-2.673	0.139	0.000	-2.508	0.266	0.000
18 years	-2.743	0.159	0.000	-3.313	0.423	0.000
19 years	-2.778	0.167	0.000	-2.836	0.348	0.000
20 years	-2.647	0.160	0.000	-2.598	0.314	0.000
21 years	-2.794	0.172	0.000	-4.132	0.618	0.000
22 years	-2.618	0.167	0.000	-2.695	0.374	0.000
23 years	-2.544	0.164	0.000	-3.104	0.477	0.000
24 years	-2.711	0.195	0.000	-3.147	0.517	0.000
25 years	-2.522	0.171	0.000	-2.625	0.393	0.000
26 years	-2.556	0.188	0.000	-2.905	0.515	0.000
27 years	-2.379	0.175	0.000	-3.975	1.002	0.000
28 years	-2.830	0.228	0.000	-2.901	0.474	0.000
29 years	-2.484	0.201	0.000	-11.314	0.096	0.000
30 years	-2.209	0.106	0.000	-2.857	0.271	0.000
Year effects (reference category 1963)						
1964	0.022	0.111	0.421	0.115	0.178	0.259
1965	0.066	0.109	0.273	-0.121	0.187	0.258
1966	-0.025	0.114	0.415	0.189	0.178	0.144
1967	-0.161	0.118	0.085	-0.367	0.198	0.032
1968	-0.112	0.110	0.153	-0.079	0.182	0.332
1969	-0.271	0.117	0.010	0.047	0.176	0.395
1970	-0.199	0.117	0.044	0.091	0.174	0.300
1971	-0.307	0.124	0.007	0.158	0.173	0.179
1972	-0.256	0.108	0.009	0.526	0.152	0.000
1973	-0.085	0.096	0.186	0.311	0.148	0.018
1974	-0.225	0.103	0.014	0.480	0.144	0.000
1975	-0.329	0.109	0.001	0.011	0.165	0.473
1976	-0.152	0.113	0.090	-0.035	0.167	0.416
1977	-0.118	0.111	0.145	-0.275	0.190	0.074

1978	-0.103	0.107	0.168	0.079	0.157	0.306
1979	-0.128	0.106	0.115	0.252	0.156	0.053
1980	-0.161	0.112	0.075	0.191	0.163	0.121
1981	-0.130	0.114	0.127	0.219	0.169	0.098
1982	0.043	0.106	0.342	0.229	0.165	0.083
1983	0.084	0.100	0.202	0.398	0.151	0.004
1984	0.005	0.103	0.482	0.494	0.148	0.000
1985	-0.061	0.109	0.286	0.445	0.158	0.002
1986	0.047	0.110	0.335	0.380	0.159	0.008
1987	0.086	0.111	0.219	0.590	0.155	0.000
1988	0.029	0.098	0.384	0.710	0.139	0.000
1989	-0.035	0.100	0.363	0.718	0.141	0.000
1990	0.117	0.103	0.128	0.323	0.154	0.018
1991	0.229	0.110	0.018	-0.071	0.173	0.341
1992	0.590	0.102	0.000	-0.147	0.181	0.207
1993	0.593	0.095	0.000	-0.110	0.165	0.253
1994	0.188	0.099	0.030	0.131	0.156	0.201
1995	0.283	0.103	0.003	0.084	0.157	0.296
1996	0.187	0.105	0.038	0.110	0.156	0.242
1997	0.048	0.101	0.317	0.503	0.149	0.000
1998	0.105	0.094	0.131	0.418	0.137	0.001
1999	0.024	0.098	0.404	0.339	0.142	0.008
2000	-0.077	0.099	0.220	0.373	0.145	0.005
2001	-0.104	0.105	0.162	0.308	0.150	0.020
2002	-0.018	0.100	0.429	0.259	0.147	0.039
2003	0.000	0.099	0.498	0.219	0.145	0.065
2004	-0.073	0.115	0.264	0.309	0.166	0.031
Age effects (reference category 23)						
24	0.024	0.057	0.334	-0.029	0.086	0.367
25	-0.162	0.066	0.007	0.118	0.097	0.111
26	-0.213	0.068	0.001	-0.100	0.104	0.169
27	-0.345	0.063	0.000	-0.210	0.088	0.009
28	-0.476	0.063	0.000	-0.126	0.081	0.061
29	-0.368	0.070	0.000	-0.175	0.093	0.030
30	-0.494	0.081	0.000	-0.081	0.106	0.220
31	-0.563	0.084	0.000	-0.191	0.110	0.042
32	-0.585	0.075	0.000	-0.150	0.096	0.060
33	-0.564	0.070	0.000	-0.200	0.093	0.016
34	-0.552	0.082	0.000	-0.106	0.101	0.148
35	-0.593	0.089	0.000	-0.196	0.116	0.045
36	-0.526	0.086	0.000	-0.413	0.126	0.001
37	-0.584	0.082	0.000	-0.230	0.114	0.022
38	-0.599	0.081	0.000	-0.414	0.110	0.000
39	-0.488	0.085	0.000	-0.302	0.117	0.005
40	-0.618	0.095	0.000	-0.300	0.128	0.010
41	-0.604	0.098	0.000	-0.345	0.135	0.005
42	-0.694	0.094	0.000	-0.307	0.122	0.006
43	-0.609	0.089	0.000	-0.312	0.113	0.003
44	-0.710	0.104	0.000	-0.381	0.142	0.004
45	-0.523	0.105	0.000	-0.456	0.157	0.002
46	-0.507	0.104	0.000	-0.476	0.164	0.002
47	-0.650	0.101	0.000	-0.463	0.143	0.001
48	-0.649	0.100	0.000	-0.489	0.150	0.001
49	-0.632	0.109	0.000	-0.649	0.168	0.000
50	-0.552	0.107	0.000	-0.782	0.190	0.000
51	-0.527	0.110	0.000	-0.982	0.221	0.000
52	-0.358	0.100	0.000	-0.783	0.185	0.000
53	-0.373	0.099	0.000	-0.666	0.194	0.000
54	0.031	0.105	0.384	-0.842	0.198	0.000

55	0.066	0.106	0.267	-0.891	0.221	0.000
56	0.131	0.112	0.122	-0.760	0.236	0.001
57	-0.047	0.108	0.334	-0.903	0.237	0.000
58	0.166	0.107	0.060	-0.849	0.250	0.000
59	0.631	0.109	0.000	-1.015	0.301	0.000
60	0.495	0.120	0.000	-1.001	0.373	0.004
61	0.718	0.123	0.000	-0.550	0.313	0.039
62	0.803	0.122	0.000	-1.025	0.369	0.003
63	0.859	0.124	0.000	-1.177	0.429	0.003
64	2.394	0.106	0.000	-1.437	0.715	0.022
Unobserved heterogeneity						
m2	-0.330	0.215	0.062	-1.915	0.139	0.000
m3	-1.034	0.171	0.000	-1.016	0.074	0.000
p2	1.737	0.413	0.000			
p3	1.465	0.275	0.000			
Pseudo log-likelihood						-4385653

Notes: Maximum likelihood estimates using sampling weights. Standard errors based on Huber-White sandwich estimator.

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